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HAZARDS OF EXPROPRIATION: TENURE INSECURITY AND INVESTMENT IN RURAL CHINA

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This paper uses household data from Northeast China to examine the link between investment and land tenure insecurity induced by China's system of village-level land reallocation. We quantify expropriation risk using a hazard analysis of individual plot tenures and incorporate the predicted "hazards of expropriation" into an empirical analysis of plot-level investment. Our focus is on organic fertilizer use, which has long lasting benefits for soil quality. Although we find that higher expropriation risk significantly reduces application of organic fertilizer, a welfare analysis shows that guaranteeing land tenure in this part of China would yield only minimal efficiency gains. (*JEL* P32, Q15)

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Secure property rights are considered an important catalyst for economic growth, the argument being that investment can only flourish when there is a reasonable chance of reaping its rewards. Reduced risk of capital expropriation by the state historically has contributed to higher growth (Douglass C. North and Barry R. Weingast, 1989). Today, insecurity of ownership remains a major issue in current and former socialist economies, and fear of expropriation may even impede privatization (Raul Laban and Holger C. Wolf, 1993). Yet, little, if any, empirical evidence exists either on how individual investors respond to heightened expropriation risk, or on the social cost of insecure property rights. A fundamental empirical limitation is that, even in contexts where the threat of expropriation is ubiquitous, expropriation itself is usually a rare event, and one that is typically observed only on a macro scale.[1]

In this paper, we examine the link between investment and expropriation risk using household-level data from rural China. Under China's current system of land management, local leaders periodically reallocate collectively-held land among farm households in the same village. Made at the village level, these reallocation—or expropriation—decisions vary across communities and are largely exogenous from the individual farmer's perspective, though the chance of being involved in a reallocation may depend on certain household characteristics (Scott Rozelle and Guo Li, 1998; Michael R. Carter, et al., 1995). Importantly, the timing of these reallocations is uncertain for the individual farmer. In this sense, rural China is an ideal case-study of tenure insecurity, providing the requisite cross-sectional variation in expropriation risk. The methodological problem, and the main innovation of this paper, is to quantify this risk and incorporate it into a tractable empirical model of investment behavior. We do so by performing a hazard analysis of individual plot tenures and relating the predicted "hazards of expropriation" to land-specific investment, specifically investment in soil quality. Because the theoretical model provides a characterization of the social cost of tenure insecurity, we can use our estimates to assess the potential welfare gains from changes in China's land management system.

To motivate the hazards approach, consider how one might proxy for tenure insecurity in a regression explaining plot-level farm investment. One strategy would be to use the length of time the farmer has held the plot, plot tenure, as such a proxy. The longer a farmer has held the plot, the more secure he must feel about keeping the plot in the future, and the more he will invest in the plot. There are three distinct problems with this strategy. First, actual plot tenure is a realization of a stochastic process, and so is

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[1] Jonathan Eaton and Mark Gersovitz (1983) make a similar point in the context of foreign investment. Using cross-country data, Philip Keefer and Stephen Knack (1997) investigate the role of aggregate measures of expropriation risk in growth regressions. Henning Bohn and Robert T. Deacon (2000) also use cross-country data to examine the impact of political variables on investment and natural resource extraction.

determined by the luck of the draw. For this reason, plot tenure is a noisy indicator of the underlying uncertainty facing the farmer; in other words, we have an errors-in-variables problem. Second, plot-level expropriation risk (as opposed to realized plot tenure) may be determined by the same unobserved variables as plot-level investment, which would lead to a simultaneity bias. For example, farmers may be more likely to lobby against or resist reallocations of their more fertile plots and as a result hold them longer. At the same time, farmers may invest more heavily in these more fertile plots. Third, there is a difficulty in interpretation. Longer plot tenure may indeed imply lower expropriation risk; alternatively, it may imply that the plot is becoming "due" for a reallocation. The question of whether expropriation risk rises or falls with plot tenure is, in technical language, a question of whether there is positive or negative duration dependence in the hazard rate of plot expropriation. Thus, the relationship between investment and tenure insecurity is intimately tied to the form of the hazard of expropriation.

A number of recent empirical studies have investigated the relationship between land rights and farm productivity or investment in various countries, but none of these studies directly examines the impact of expropriation risk.[2] For example, since land expropriation is not important in rural Ghana, Besley's (1995) analysis of farm investment focuses on the impact of land transfer rights. Similarly, the work of Feder and collaborators on the effect of land titles in rural Thailand does not directly address expropriation risk. Because Thai farmers face little threat of eviction, the primary benefit of legal title is in enabling households to access credit by using land as collateral. In China, banks, by law, cannot take land as collateral and policy explicitly prohibits land sales, so our analysis puts the spotlight directly on the impact of expropriation risk.

The plan of the paper is as follows. In the next section, we provide the institutional background on China's land management system, and present some evidence strongly suggesting a link between property rights and investment in soil quality. Section II lays out a theoretical model of farm investment in the presence of expropriation risk and derives its empirical implications. In section III, we discuss the empirical implementation of the hazard model and section IV presents the hazard model results. Sections V and VI discuss, respectively, the empirical strategy and results of the fertilizer demand estimation. These results form the basis for our quantitative analysis of agricultural policy reform in section VII. Section VIII concludes the paper.

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[2] A partial listing includes Gershon Feder and Tongroj Onchan (1987), Yongyuth Chalamwong and Gershon Feder (1988), Frank Place and Peter Hazell (1993), Timothy Besley (1995), and Lee J. Alston, et al. (1996).

I. Institutional Background and Data

A. Land Rights and Tenure in China

While the Household Responsibility System, instituted in the early 1980's, gave farmers clear rights to the residual income from farm activities, reformers kept local officials firmly in control of the allocation and general management of land resources. Most cultivated land in rural China, with the exception of a small fraction that is still managed as state farms, remains collectively owned; none is truly privatized. National policy proclamations provide *guidance* to regional and community leaders, encouraging the extension of long-term lease rights and unconstrained transfer rights. The state, however, has purposely provided localities flexibility in land management and has explicitly allowed leaders the right to make periodic adjustments to household land holdings if conditions so require (James K. Kung and Shouying Liu, 1997).

Why do local leaders reallocate land, especially if they know that there is a potentially adverse impact on farmer investment incentives? Although there is great heterogeneity across villages, the empirical literature has uncovered three main motivations. First, reallocations help maintain the egalitarian distribution of land in the face of household-level demographic change, including new household formation (James K. Kung, 1994). Second, reallocations help eliminate the growing inefficiency in the distribution of land across households caused by demographic changes in conjunction with poorly functioning land rental and labor markets (Guo Li, 1999; Dwayne Benjamin and Loren Brandt, 2000). Third, village leaders use land reallocations as a "carrot and stick" to fulfill output quotas and collect taxes (Rozelle and Li, 1998).

Land in most villages can be divided into two types, private plots (*ziliu di*) and collectively controlled land (*jiti di*), although collective land predominates (Yuk-shing Cheng and Shuki Tsang, 1995). While in theory the collective still ultimately *owns* the private plots and farmers cannot sell them, most village leaders do not intervene in decisions on private plots. Farmers nearly always have rights to residual output, unfettered rental rights, and enjoy a high degree of tenure security. Collectively controlled land is of three nationally recognized types: ration land (*kouliang tian*), granted to farmers to meet household subsistence requirements; responsibility land (*zeren tian*), granted to farmers on the condition that they deliver low-priced grain and cotton quotas to the state; and contract land (*chengbao tian*), which village leaders lease to farmers for a fee, but for an uncertain duration (Heng Liang, 1993).

A 1996 village level survey shows that private plots are held on average 12 years longer than collective plots, evidence of the relative insecurity of tenure on collectively controlled land.[3]

B. Land Rights and Soil Quality Investments

Our focus is on investment in soil quality through the application of organic fertilizer, a mixture of manure, dredged soil, decayed vegetable matter, and other farm-yard wastes. Farmers can spend days collecting, mixing, transporting, and incorporating organic fertilizer into the soil (Qiaolun Ye and Scott Rozelle, 1994). The average farm household in our sample allocates about 8 percent of its annual labor days spent on cultivation to organic fertilizer application, making it one of the most time intensive tasks. Although organic fertilizer contains trace amounts of nitrogen and other minerals that promote healthy crop growth, its primary benefit is in maintaining soil structure, particularly the ability of the soil to retain moisture (China, Ministry of Agriculture, 1984). This benefit is long-lasting; a single application of organic fertilizer in most sub-tropical and temperate climatic zones (areas covering most of China and all of the sample locations) can have an effect on the soil for four to five years. In contrast, the effects of chemical fertilizers, principally nitrogen and phosphate, last only for a single growing season.

To be sure, there are fixed investments in cropland, such as well-digging, surface irrigation, drainage, and terracing, that our analysis ignores.[4] In the case of rural China, however, limiting our focus to recurrent investment in soil quality is justified for two reasons. First, at the household level, many of these fixed investments are rare during the reform era; recent analyses of retrospective data in China show that only around one percent of households in a given year undertake such activities either on their own initiative or at the direction of the village (Chengfang Liu, 2001; Linxiu Zhang, 2001). This paucity of investment in farm infrastructure over the past two decades is likely to be due much less to poor land rights (as we discuss next) than to the fact that so much investment was undertaken during the proagriculture days of the prereform era.

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[3] The survey of 215 randomly selected villages in six provinces also shows that the average length of tenure on responsibility plots is 7 years, on ration plots is 5 years, and on contract plots is 4 years. Rental and control rights also vary by land tenure type.

[4] We also ignore tree-planting (which Besley, 1995, finds to be responsive to land-rights in Ghana) and, more generally, transformation of cropland into orchards, fishponds, and so forth. Our data are not suited for this type of investigation. But, in any case, the scope for such investments, even under full private property rights, would be limited in China (which currently has around 5 percent of cropped area in perennials, already a fairly high amount), just as it is in the agricultural economies of other large continental nations, such as the United States (1 percent), Europe (2.5 percent), or India (2.2 percent).

Second, many fixed investments either do not depend directly on rights to a specific piece of cultivated land, or are more efficiently organized at the communal or village level irrespective of the property rights regime. For example, wells in China are almost always either located between plots or on the fringes of roads and paths and not in the middle of fields. Hence, reallocation of a plot does not usually involve expropriation of a well that sits on it. Investments such as canal irrigation, drainage, and terracing projects, as Besley (1995) points out in the context of Ghana, often affect the land of several farmers. To the extent that such investments create public goods, and also have high capital costs, the community, rather than the household, may be the most natural unit to undertake them, since the community can internalize the externalities and mobilize the necessary resources. In China, then, we argue that soil quality improvement, while perhaps not the only investment in cropland responsive to plot-specific rights, is certainly one of the most important ones.

C. Data and Preliminary Evidence

This paper uses data from a World Bank Living Standards Measurement Survey of 727 farm households from 31 villages in Hebei and Liaoning Provinces conducted in the summer of 1995. Hebei and Liaoning, located in the northeast, are two of China's major agricultural provinces. The survey collected detailed information on household characteristics, agricultural production, and non-farm activities. There was also a comprehensive survey of village leaders. Appendix Table A.1 describes the main samples used in this paper. Total landholdings of each household were recorded on a plot-by-plot basis. Farmers were asked about the characteristics of each of the 3,113 plots, of which 2,898 are collectively controlled and the remainder are private. A special feature of the survey is a plot comparison module. Two plots from each household with different tenure type (e.g., one private plot and one plot from one of the categories of collective land) were selected for more detailed investigation, including the comprehensive enumeration of plot-specific inputs and outputs. This module yielded a sample of 1074 plots from 612 households growing the four main crops (maize, rice, cotton, and soybean); 961 of these plots are collectively controlled.

To provide some initial evidence on the importance of land rights for investment behavior, we compare fertilizer use, organic and chemical, on private and collectively controlled plots (see section VI for precise definitions of the fertilizer variables). Recall that private plots are generally held much longer than collective plots and are considered much more secure. Table 1 presents household fixed effects estimates of plot-level fertilizer intensity as a function of plot characteristics. The sample of 216 plots is based on the 108 households that report both a private and a collective plot in the plot comparison module. By comparing plots cultivated by the *same* household, we control for household and village-level unobservables that may influence fertilizer use. Among the plot characteristics is a dummy variable for whether the plot is private or collectively controlled.[5] We use the fixed effect tobit estimator of Bo E. Honoré (1992) to account for censoring, most severe in the case of phosphate. The results show that organic fertilizer use is significantly higher on private plots, but that this is not the case for chemical fertilizers. Besides freedom from expropriation, however, there may be other rights enjoyed on private plots and not on collectively controlled plots that lead to greater organic fertilizer use. Thus, the evidence in support of the tenure security hypothesis can only be viewed as circumstantial at this point. Later, we assess the findings in Table 1 in light of our estimates of the structural model.

II. Theoretical Framework

A. An Investment Model with Expropriation Risk

We begin with a standard investment model, augmented to allow for the risk of land expropriation. From the farmer's perspective, village leaders will expropriate one of his plots at some random time τ. Let the survivor function $S(t) = \Pr(\tau \geq t)$ denote the probability that the farmer retains control of the plot until at least time *t*. The corresponding hazard function of expropriation is defined as $h(t) = -\dot{S}(t)/S(t) = f(t)/S(t)$, where $f(t)$ is the probability density function of plot tenure. The hazard function represents the instantaneous probability of losing one's plot, given that it has already been held for *t* years.

Our focus is on investments in soil quality on a given plot of land. All choice variables are normalized by plot area. Denote the stock of organic matter in the soil at time *t* by $k(t)$. Farmers may apply a quantity of chemical fertilizer, $n(t)$, to the plot at a cost of $c_n n(t)$, but doing so enhances only current period output. A quantity of organic fertilizer, $x(t)$, may also be applied to the plot at a cost of $c_x x(t)$, which raises the stock of organic matter in the soil currently and into the future. Normalizing

[5] We do not control for irrigation since this varies across private and collective plots within only 4 of the households.   A referee suggested that the "quality" of irrigation may nevertheless vary by tenure type, but this seems unlikely as our data show more than three-quarters of the households with irrigated plots use groundwater. With groundwater irrigation and the plastic hose technology common in this part of China, access to water should not differ across plots as might be the case with canal irrigation. We also do not control for the crop grown on the plot for reasons discussed in section VI. In any case, all but 11 of the 216 plots grew maize in the fall season, and adding a maize dummy has virtually no impact on the estimates in table 1.

the price of output to unity, the value of output from the plot is given by the neoclassical production function $F(n(t),k(t);\varphi)$, where φ is an index of plot fertility. For simplicity, this production function ignores other inputs, an issue we return to below, as well as multiple crops, which we discuss in section VI.

Farmers choose the time-path of fertilizer application to maximize expected discounted net returns on the plot. Thus, letting $R(t)=F(n(t),k(t);\varphi)-c_n n(t)-c_x x(t)$ be the net yield and *r* be the discount rate, the farmer's problem is

(1) $$Max\int_0^{\infty} e^{-rt} S(t)R(t)dt$$

subject to

(2) $$\dot{k}(t)=-\delta k(t)+x(t)$$

(3) $$x(t)\geq 0,\ n(t)\geq 0$$

and $k(0)=k_0$. The farmer's objective function (1) reflects the fact that he collects the period *t* return on investment only with probability $S(t)$. According to (2), the stock of organic matter is depleted at the fixed rate δ. Farmers may also choose not to replenish this stock at all (i.e., $x(t)=0$).

The following equations, derived in Appendix A, describe the optimal choice of fertilizers in this model in the case where $x(t)>0$ and $n(t)>0$:

(4) $$F_k(t)=(r+\delta+h(t))c_x$$

(5) $$F_n(t)=c_n$$

(6) $$x(t)=\delta k(t)+\frac{c_x F_{nn}(t)}{\Omega(t)}\dot{h}(t),$$

where $\Omega=F_{nn}F_{kk}-F_{nk}^2>0$ by the concavity of *F*. Using only equations (4) and (5), we can solve for the unconditional chemical fertilizer demand $n^*(t)=n^*(h(t),r,\delta,c_x,c_n,\varphi)$ as well as for $k^*(t)$. Combining these results with equation (6) yields the organic fertilizer demand function, $x^*(t)=x^*(h(t),\dot{h}(t),r,\delta,c_x,c_n,\varphi)$.

Organic fertilizer use depends not only on the hazard of expropriation, $h(t)$, but also on how the hazard changes with plot tenure--i.e., on duration dependence $\dot{h}(t)$. In other words, it matters whether village leaders are more or less prone to reallocate land as time elapses since the last reallocation. Recall that $\dot{h}(t)>0$ means positive duration dependence (expropriation risk increases over time), while

$\dot{h}(t) < 0$ means negative duration dependence. Consider the case of zero duration dependence. According to equation (6), when $\dot{h}(t) = 0$, farmers will invest just enough to maintain the stock of organic matter in the soil. However, if $\dot{h}(t) > 0$, farmers will invest less than necessary to cover soil depletion, since $c_x F_{nn} / \Omega < 0$. Intuitively, the shadow price of capital, on the right-hand side of equation (4), is rising with length of tenure on the plot. Conversely, if $\dot{h}(t) < 0$, farmers will apply more organic fertilizer than necessary to cover soil depletion because the shadow price of capital is falling with time. By contrast, the demand for chemical fertilizer, $n^*(t)$, is not *directly* influenced by $\dot{h}(t)$. Since this type of fertilizer is not an investment, behavior is myopic—choices are made without heed to future changes in the shadow price. However, $n^*(t)$ does depend on $h(t)$ through a cross-price effect.

Comparative statics are readily obtained by assuming the proportional hazards form $h(t;\theta) = g(t)\theta$, where $g(t)$ is a pure function of time and θ can be viewed as the exogenously given frequency of expropriation. Note that $\dot{h}(t;\theta) = h(t;\theta)\,\dot{g}(t)/g(t)$. Also, assume that all third derivatives of *F* are zero (or at least are negligible); in other words, *F* is approximately quadratic. To understand how an exogenous change in policy governing tenure security would affect fertilizer use, we differentiate equations (4) to (6) with respect to θ, which yields

(7) $$\frac{dx^*(t)}{d\theta} = \frac{c_x F_{nn}}{\Omega}\left[\delta g(t) + \dot{g}(t)\right]$$

(8) $$\frac{dn^*(t)}{d\theta} = -\frac{c_x F_{nk} g(t)}{\Omega}$$

The necessary and sufficient condition for a negative relationship between investment and expropriation risk is that $\dot{g}(t)/g(t) > -\delta$;[6] in other words, the hazard cannot fall at a proportional rate faster than the depreciation rate. Intuitively, an increase in θ has two effects. First, it increases the hazard rate, which increases the shadow price of capital directly. This effect lowers the optimal stock of capital and so decreases the amount of investment necessary to cover depreciation. The second effect of the rise in θ is to increase the absolute value of the *slope* of the hazard function at each *t*. The impact on investment depends on whether this slope is positive or negative. When the slope is negative, i.e., if expropriation becomes increasingly rare as plots are held longer, the rise in θ accentuates the decline in the shadow price of capital over time, and could thus result in a net increase in the investment rate. However, during any period in which the hazard of expropriation is increasing over time, investment is unambiguously decreasing in the frequency of expropriation; i.e., $\dot{g}(t) > 0 \Rightarrow dx^*(t)/d\theta < 0$.

Use of chemical fertilizer is influenced by changes in θ indirectly, through a cross-price effect. For example, if chemical fertilizer is a substitute for organic fertilizer ($F_{nk} < 0$), then $dn^*(t)/d\theta > 0$. However, the situation becomes complicated when there are two chemical fertilizers (as is the case in our data), or when there are multiple inputs more generally. With a second chemical fertilizer, it is no longer sufficient that $F_{nk} < 0$ for $dn^*(t)/d\theta > 0$; the result also depends on the substitutability between the second chemical and organic fertilizer *and* on the substitutability between the two types of chemical fertilizers as well.[7]

B. The Social Cost of Expropriation Risk

To quantify the efficiency cost of expropriation risk, we must define the total social return or value of the representative plot.[8] The social value exceeds the private value of a given plot for the simple reason that when the plot is expropriated from one farmer it does not cease to exist, but rather is transferred to another farmer. To take such transfers into account, define the social value function *V* at date 0 as the solution to

(9) $$V(k(0)) = \int_0^\infty e^{-rt}\left[S(t)R^*(t) + f(t)V(k(t))\right]dt$$

The first part of (9) is the maximized value of the individual farmer's objective function, or just the private value of the plot. The additional term reflects the fact that if this farmer loses the plot at date *t*, which occurs with probability $f(t) = S(t)h(t)$, it reverts to another farmer in whose hands the social

[6] Using the Weibull form $g(t) = \alpha t^{\alpha-1}$ to be discussed in section III, the condition becomes $\alpha > 1 - \delta t$, which is more likely to be satisfied as time elapses.

[7] As alluded to earlier, the technological relationship between chemical and organic fertilizers is complex. While manure provides trace amounts of nitrogen, its main benefit is improved soil structure, which can allow more effective use of inorganic nitrogen and other chemical fertilizer. Thus, it is an open question whether, on balance, organic and chemical fertilizer are substitutes or complements in production.

[8] We ignore the portfolio risk induced by land expropriation. Presumably, if farmers are risk averse, they prefer more certain plot tenure, holding constant expected tenure. Estimating the utility cost of uninsured expropriation risk is beyond the scope of this paper.

value of the plot is again *V*, but discounted back from date *t*. Notice that, in general, *V* is nonstationary, since the stock of organic matter in the soil changes over time. However, if $k(t)$, and hence *V*, is approximately stationary, then we may rearrange terms in (9) and integrate by parts to get

$$V = \frac{\int_0^\infty e^{-rt} S(t) R^*(t) dt}{r \int_0^\infty e^{-rt} S(t) dt} \tag{10}$$

The constant hazard form $h(t) = \theta$ delivers explicit analytical results because in this case not only is *V* stationary but $R^*(t)$ is just equal to the constant $R^* = F(n^*, k^*; \varphi) - c_n n^* - c_x x^*$. Equation (10) then reduces to $V = R^* / r$. By contrast, under these same assumptions, the private value of the plot is $R^* / (r + \theta) < V$. Furthermore, $\partial R^* / \partial \theta = c_x [(r + \theta)/\delta](\partial x^* / \partial \theta)$ by the envelope theorem, so the marginal impact of expropriation risk on net yield is proportional to its marginal impact on the investment input; $\partial R^* / \partial \theta$ vanishes as the depreciation rate approaches infinity, since organic fertilizer ceases to be an investment. Substituting for c_x using (4) gives the marginal social cost of expropriation risk in elasticity form

$$\eta_\theta \equiv \frac{\partial V}{\partial \theta} \frac{\theta}{V} = \frac{r + \theta}{r + \delta + \theta} \frac{F_x x^*}{R^*} \varepsilon_\theta , \tag{11}$$

where $F_x = F_k / \delta$ and $\varepsilon_\theta = (\partial x^* / \partial \theta)(\theta / x^*)$ is the elasticity of investment with respect to expropriation risk. Taking the three terms in this expression in reverse order, the marginal social cost of tenure insecurity depends *directly* on the responsiveness of investors to expropriation risk and on the output contribution of organic fertilizer investment relative to net yield, and *inversely* on the rapidity of depreciation. That η_θ is also increasing in *r* is due to the fact that a higher *r* leads to a lower stock of organic matter and therefore to a higher marginal product of organic fertilizer. As a result, increases in the risk of expropriation become more costly at the margin.

In our quantitative assessment of agricultural policy reform in China, presented in Section VII, we use a generalization of the above formula that takes duration dependence into account.

III. Estimation Strategy: Hazard Analysis

The econometric analysis proceeds in two steps: (1) estimation of a hazard model of length of tenure on each plot i to get a plot-specific prediction of expropriation risk, $\hat{\theta}_i$; and (2) estimation of the fertilizer demands as functions of $\hat{\theta}_i$. We discuss the first step of the analysis in this and the next section, and the second step in sections V and VI.

The expropriation risk parameter is specified as

(12) $$\theta_i = \exp(\beta' X_i + \zeta_v)$$

where X_i is a vector of plot and, possibly, household characteristics, and ζ_v is a village-specific intercept estimated using village dummy variables. The inclusion of village dummies here is motivated by the observation that the land policy environment, particularly the frequency of reallocations, not only differs across villages but is also stable within villages over time. Evidence on the latter point is taken from the village survey, which shows that the most important decision-makers in the 31 villages of our sample, the party secretary and village leader, have typically been in power (or at least participated in the management of village affairs) since before the start of the household responsibility system.[9] Correspondingly, the survey indicates that the main village land policies have also barely changed between 1995 and 1988, the earliest year for which such information is available.[10] Given that the village leadership decides to make an adjustment, certain types of plots may be more prone to reallocation, hence the inclusion of plot characteristics in X_i. Finally, the chance of losing a plot may depend on the household's demographic characteristics, since village leaders supposedly care, at least partly, about equalizing landholdings across households.

Our goal is to estimate the distribution of times between plot reallocations, or, equivalently, the hazard function $h(t;\theta)$. However, the data only tell us how long the farmer has held a plot, not how long he *will* hold the plot. In the parlance of duration models, the first type of data (the data that we have) is an incomplete or right-censored duration; the second type is a complete duration. Unless the

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[9] For example, the party secretary, the key decision-maker in many, if not most, villages, has been in a leadership position in the village for an average of 19 years (median 20 years) as of 1995. Similarly, the average tenure of the village leader, the other important village-level official, is 17 years (median 17 years).

[10] Specifically, in 98 percent of the villages, leaders report that the policy on land adjustments was the same in 1995 as it was in 1988. Likewise, in 96 percent of the villages, the right to rent out land has not changed, and, in 95 percent of villages, the right of the farmer to choose his own crops has not changed.

duration distribution is exponential, the distribution of incomplete durations will differ from the distribution of complete durations. Denote the complete duration distribution by $f(t_i;\theta_i)$, where *i* indexes plots. We may write $f(t_i;\theta_i) = h(t_i;\theta_i)S(t_i;\theta_i)$. By contrast, the distribution of incomplete durations is (according to D. R. Cox, 1970)

(13) $$f^*(t_i^*;\theta_i) = \frac{S(t_i^*;\theta_i)}{E(t_i)} = \frac{S(t_i^*;\theta_i)}{\int_0^\infty S(\tau;\theta_i)d\tau}$$

where t_i^* is the length of time that plot *i* has been held and *E* is the expectations operator. This density is the individual contribution to the likelihood function that we maximize.

Our empirical specification of the baseline hazard takes the standard Weibull form, $g(t) = \alpha t^{\alpha-1}$, so that the survivor function is

(14) $$S(t_i^*;\theta_i) = e^{-\theta_i (t_i^*)^\alpha}$$

where α is the duration dependence parameter and $\alpha > 1$ implies positive duration dependence. An advantage of the Weibull is that the denominator of (13) integrates to the gamma function so that numerical integration is unnecessary in calculating the likelihood function.

There are two additional issues to consider in the specification of the hazard model. First, because decollectivization occurred in the early 1980's, plot tenure durations are left-censored at the time since the reform, t_R. That is, a farmer living in a village that decollectivized in, say, 1982 who was interviewed in 1995 could not have held on to his collectively controlled plot for more than 14 years. Further complicating matters is the fact that the decollectivization process began at different times in different villages and continued over a period of months or years. As a result, part of the variation in plot tenure is due to variation across villages in the timing of the reform and not in the underlying risk of expropriation. A tractable approach to dealing with this problem is to set a fixed censoring point, $t_R = 12$, for all observations, corresponding to the end of the reform period in 1984. Thus, we define a censoring indicator d_i, which equals one if $t_i \geq t_R$ and zero otherwise. The resulting likelihood function, shown in Appendix B, resembles that of a tobit model.

Another general source of bias in hazard models is the presence of unobserved heterogeneity; an example of which in this case is unobserved plot fertility, φ. Misspecification of the unobserved heterogeneity distribution can lead to biases in the predicted $\theta's$, which in turn may bias estimates of the fertilizer demand functions. To address this issue, we estimate our hazard model using the nonparametric maximum likelihood technique of James J. Heckman and Burton Singer (1984). Appendix B details the implementation of this approach in the present context.

The end product of the hazard analysis is a predicted value of $\log\theta_i$ for each plot; i.e., $\log\hat{\theta}_i = \hat{\beta}'X_i + \hat{\zeta}_v$. To gain intuition for the hazard rate approach, consider the case where $\beta = 0$. Conditional on α, it can be shown that the maximum likelihood estimate (MLE) of $\log\theta_i = \zeta_v$ is

(15) $$\hat{\zeta}_v(\alpha) = -\log(\alpha) - \log\left(\frac{1}{N_v}\sum_{i=1}^{N_v}(t_i^*)^\alpha\right)$$

where N_v is the number of plots in village *v*. A second implicit equation, suppressed here for brevity, can be used to solve for $\hat{\alpha}$ and $\hat{\zeta}_v$. In the absence of duration dependence ($\alpha = 1$), equation (15) collapses to $\hat{\zeta}_v = -\log(\bar{t}_v^*)$, or minus the log of the village average plot tenure. This is so because when the distribution of complete durations is exponential, so is the distribution of incomplete durations. Thus, the inverse of the average incomplete duration is the MLE of θ. In the presence of duration dependence, the MLE of θ (both conditional and unconditional on α) is a complicated function of t_i^*, depending on the whole distribution of durations rather than on just the mean.

IV. Results of Hazard Analysis

Table 2 shows the frequencies of plot tenure durations for the three categories of collective land (tenure type) in the full sample of 2,898 collectively controlled plots. Indicative of the left-censoring noted above is the considerable "mounding" of observations at durations of 12 and 13 years, corresponding to the end of decollectivization period.[11]

Regressors for the hazard analysis include plot characteristics—tenure type, area, self-reported land quality (collapsed into two categories: high and low), whether irrigated or not, and topography (flat

[11] Another type of left-censoring, which could potentially bias the hazard estimates, is due to farms splitting up. If a son inherited his father's collective plot, say five years ago, then the reported tenure on the plot is five years even though his father may have acquired the plot in a reallocation more than five years ago. The duration of interest is the time since reallocation, not the time since inheritance. Unfortunately, we do not have information on how each plot was acquired. We do, however, have such information at the household level, according to which only 3 percent of recorded land acquisitions involved the establishment of a new household. Thus, this type of left-censoring is unlikely to impact our hazard estimates.

versus hilly or terraced)—as well as household size and composition. We only have household demographic information at the time of the survey, not at the time of plot acquisition and certainly not over the entire duration of tenure on the plot, an issue we return to below.[12] To explore whether these demographic variables influence the reallocation decisions of village leaders, as well as other specification issues, we first estimate a hazard model without village dummies and without controlling for unobserved heterogeneity. The first thing to notice in specification (1) of Table 3 is that the duration dependence parameter, $\log \alpha$, is statistically indistinguishable from zero, which means that there is zero duration dependence. The reported t-values in specification (1) account for intra-village clustering.

The coefficient estimate for household size in specification (1) indicates that expropriation risk is lower on plots held by larger households. This finding suggests that village leaders are motivated, in part, by equity considerations, allowing larger households to keep their plots longer on average than smaller households. A referee suggested a second effect also may be at work. Suppose that plots of smaller households are indeed expropriated and transferred to larger households in order to equalize per capita landholdings in the village. We might find that in villages that have just had such a reallocation, plots held by larger households tend to have short tenure durations because many of them have been newly allocated to the households. By contrast, plots held by smaller households would tend to have long durations, because their plots "survived" the reallocation in the hands of the original household. If this effect is sufficiently strong, the hazard rate would be *increasing*, rather than decreasing, in household size. Our estimates show that the first effect dominates the second.

The second specification in Table 3 uses the nonparametric maximum likelihood estimator detailed in Appendix B to control for unobserved heterogeneity. We still do not include village effects. In this case, the data support a three-point heterogeneity distribution $(M = 3)$. Most of the coefficient estimates are not substantively affected by controlling for heterogeneity. Compared to specification (1), however, the estimated duration dependence parameter increases significantly. Some increase in α would have been expected given that unobserved heterogeneity tends to bias duration dependence downward. However, the estimate of 6.8 ($e^{1.92}$) seems unreasonably large and may reflect the general difficulty in distinguishing between duration dependence and unobserved heterogeneity in hazard models (see, e.g., Heckman and Singer, 1984).

[12] For this reason, we do not include total household landholdings in the hazard model. The problem is that we do not observe landholdings *prior* to the acquisition of a given plot, only current landholdings, which may be endogenous because households that have just been allocated a plot (and hence have a short plot tenure) will tend to have more land in total.

Specification (3) includes the full set of thirty village dummies, but no unobserved heterogeneity controls. The important finding here is that the village dummies are highly jointly significant. Thus, it would appear that villages vary considerably in the frequency of reallocations.[13] Moreover, in contrast to specification (1), the estimate of $\log(\alpha)$ is significantly greater than zero, indicating positive but not, as in specification (2), unreasonably large duration dependence. When we attempt to control for unobserved heterogeneity along with the village dummies, we find that the data fail to support even a two-point distribution, suggesting that the unobserved heterogeneity detected in specification (2) might largely be a village-level phenomenon, captured by the village dummies.[14]

Finally, we return to the problem of missing data on time-varying regressors. Plot characteristics are essentially fixed during a household's tenure, as are village land policies (see discussion in section III).[15] The same cannot be said for household demographics, which change over time. The erroneous assumption that these regressors are fixed at their current values may bias the estimates of the hazard model. One option is to omit these variables from the model, though, given that our main objective is to *predict* expropriation risk, this is not necessarily the best approach. In any event, the three household demographic variables are not jointly significant in specification (3), and we obtain very similar results when we omit them in specification (4).[16]

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[13] As a data check, we compare the estimated village effects against the number of village-wide reallocations since decollectivization, as reported in the village questionnaires. Reassuringly, the correlation coefficient between the estimates, $\hat{\zeta}_v$, and the frequency of village-wide reallocations is 0.37, which is significant at the five percent level. It should be kept in mind, however, that because of frequent small-scale reallocations of land, the frequency of "village-wide reallocations " reported by village leaders likely understates the expropriation risk faced by the average farmer in a village, as reflected by the $\hat{\zeta}_v$.

[14] Specifically, in the nonparametric maximum likelihood estimation with village fixed effects, the duration dependence parameter becomes extremely large and the maximization routine fails to converge. We also tried imposing normality of the heterogeneity distribution, which gave the same result, indicating a serious identification problem.

[15] In some villages, small amounts of contract land may have been converted into responsibility land, or vice-versa, during the last 20 years. Such changes, however, almost always occur during a plot reallocation rather than within the tenure of a given household.

[16] Age of the head also changes over time, but in a deterministic fashion. While this should not affect the predictive power of the hazard model, age of the head may partly be picking up the fact that older heads have simply had more time to hold a given plot. To check for this, we reestimated specification (4) using only the 1,737 plots farmed by household heads older than forty. These heads were in their mid-twenties and would have already been managing their own farms at the time of the reform. Thus, they all could potentially have held on to their plots for the 12-year maximum duration. The coefficient on age of head

Using the estimates of specification (4), we obtain a predicted expropriation hazard rate, $\hat{h}_i(t_i) = \hat{\alpha}\hat{\theta}_i t_i^{\hat{\alpha}-1}$, for each of the 2,898 plots in the sample (specification (3) yields almost identical predictions). The median hazard rate is 0.106 (mean=0.161); in other words, at the time of the survey, the median plot had an 11 percent chance of being expropriated, given the length of time it had already been held. A better picture of expropriation risk can be obtained by calculating the probability of losing a plot in a given period of time, t_0, which is just $1-\hat{S}_i(t_0)$. Figures 1(a) and 1(b) plot these cumulative probabilities for a t_0 of five and ten years, respectively. Five years after reallocation, two-thirds of the plots have cumulative probabilities below 50 percent of being expropriated. However, 10 years after reallocation, only about a third of the cumulative expropriation probabilities are below this threshold. Indeed, about 20 percent of all plots are virtually certain (greater than 95 percent chance) of being expropriated within ten years.

V. Estimation Strategy: Analysis of Fertilizer Demand

This section considers the empirical specification of the fertilizer demand functions derived from equations (4) to (6). We focus here on the general econometric issues, leaving the detailed description of variables, sample and results for section VI.

A. Endogeneity of Expropriation Risk

When there is duration dependence, the fertilizer demand functions, $x^*(t)$ and $n^*(t)$, depend on actual plot tenure *t* (as well as on θ) through the hazard function. However, as discussed in the Introduction, introducing actual tenure into the demand functions leads to an errors-in-variables problem. We approximate the theory by assuming that, conditional on θ, fertilizer use is constant over the range of *t* in our data. Thus, we write $x^* = x^*(\theta, r, \delta, c_x, c_n, \varphi)$, and similarly for n^*. This specification puts the focus on the effect of permanent differences in expropriation risk rather than on the effect of changes

in this case is –0.0163 (2.28) and the duration parameter $\log(\alpha)$ is estimated to be 0.227 (2.95), neither of which is significantly different from its counterpart in specification (4). In sum, it appears that the coefficient on age of head, rather than confounding age and duration effects, is capturing the favorable treatment that older, more experienced, farmers receive from village leaders in land reallocations.

in expropriation risk over time. We also assume that the fertilizer demand functions take a semi-log form (in θ) so that $(\partial x^* / \partial\theta)\theta$ is constant.

In the fertilizer use analysis, as already mentioned, we use the predicted value of θ rather than its unobserved, actual value. *Actual* expropriation risk may be endogenous if, for example, village leaders use land reallocations as a "carrot and stick" to enforce village policies. In this case, the underlying risk of expropriation depends on individual household actions (e.g., farming effort) that are unobserved by the econometrician, but may be correlated with fertilizer use. Similarly, when expropriation risk depends on household efforts to lobby village leaders, more powerful farmers may face lower risk and get more fertile plots as well.[17] To the extent that the instruments in equation (12), the "first-stage regression" for expropriation risk, are orthogonal to the unobservable determinants of plot-level fertilizer use, such as an individual's farming effort and a plot's fertility, predicted expropriation risk does not suffer this endogeneity problem.

Our identification strategy relies on exogenous inter-village variation in land management policy; specifically, we include village dummies in equation (12) but not in the fertilizer demand equations directly. Indeed, the relevance of these instruments is confirmed by the very high significance of the village dummies in the hazard model. With this approach there is a danger, though, that *village* level unobservables may lead to a spurious correlation between fertilizer use and *predicted* expropriation risk. The argument involves three parts: (1) The average marginal product of organic fertilizer varies across villages in an unobservable way because of inter-village differences in plot fertility or other productive factors such as infrastructure (represented by the village mean of φ); (2) Organic fertilizer is used more intensively in villages where its average marginal product is higher; (3) Village leaders are aware that frequent land reallocation is costly because it leads to lower organic fertilizer investment, and that this cost is greater the more intensively organic fertilizer is used in the village in the first place. So, for example, in the case where $F_{k\varphi} > 0$ (i.e., the return to organic fertilizer is increasing in fertility), villages with a high φ will have both intensive fertilizer use and fewer land reallocations. In other words, rather than a high θ causing low fertilizer use, low fertilizer use may cause a high θ! Precisely the same argument applies if $F_{k\varphi} < 0$, except that in this case villages with a high φ will have both low fertilizer use and more reallocations. One way to assess the importance of this endogeneity problem is to find a

[17] Unfortunately, the household survey did not collect information on communist party affiliation or on the household's relationship to village leaders, which might have shed light on the role of political power in expropriation decisions.

proxy for village level productive endowments and include it in the fertilizer demand equations. In Section VI, we construct such a proxy using a yield production function.

B. Estimated versus Perceived Expropriation Risk

Another empirical concern is whether risk of land expropriation as estimated from the hazard model is the same as the risk that farmers *perceive* when making input decisions. If farmers have superior information to the econometrician (i.e., they have more *relevant* plot-specific knowledge than is captured by X_i), then perceived expropriation risk may diverge from predicted risk. While we cannot address this issue with the data at hand, we can ask a related question: How uncertain is plot tenure? In the extreme case, in which farmers know the exact date of eviction from their plot, there is no underlying expropriation risk even though sample variation in tenure duration may provide a sensible estimate of θ.

One way to assess whether farmers are aware of the eviction dates from their collectively-held plots is to ask them. Indeed, in the plot comparison module, the survey asked the farmer for each plot, "Do you know when your contract expires (or when it already has expired)?" If the farmer answers, "Yes," a follow up question asks: "In what year does the contract expire?" For only 32 percent of the plots (303 of the 961 collectively-held plots--there are 16 missing values), did the farmer reply that he knew when the contract would (had) expire(d). For the other 68 percent of the plots, the farmer had no idea when the contract would expire. Even among the 303 plots with known contract expiration dates, 139 (46 percent) had reported dates on or before the end of 1995. Since the survey was conducted in the summer of 1995, farmers of many of these plots probably had only just been informed of an imminent readjustment and, hence, that their contracts had expired or would expire by the end of the year. Prior to that time, including the 1994 crop year for which our fertilizer data were collected, these farmers were most likely uncertain about their future tenure. We are left then with 164 plots (54 percent of the 303 plots) for which farmers claim to know that their contract expires some time after 1995. In almost half these cases, the expiration date occurred on or before 1999. Even if farmers behave as if they are certain about their expiration date,[18] they still face tenure insecurity. In particular, once the contract expires, the plot is subject to reallocation and the farmer may lose it at some future date. Tenure uncertainty, in this case, is not eliminated but merely postponed. To sum up, our data indicate a lot of uncertainty about plot tenure,

[18] This is a strong presumption, since farmers may not believe announcements of village leaders regarding their plans to reallocate land. To the extent that they do not believe leaders, there is even more insecurity.

though for about 17 percent of the plots farmers appear to have been granted some degree of security. In the empirical work, we explore whether farmers behave (i.e., invest) any differently on this subset of plots.

C. Econometric methods

Three specific econometric issues arise in the fertilizer demand estimation: (i) censoring of fertilizer use, particularly of organic and phosphate, at zero; (ii) the presence of a regressor, $\log\hat{\theta}_i$, that has been generated from a previous estimation step and that therefore is subject to sampling error; (iii) potentially high intra-village correlation in the unobservables, or clustered data. Given that some of the explanatory variables vary mainly or exclusively between the 31 villages, including $\log\hat{\theta}_i$ (see below), failing to account for this village-level clustering could seriously understate the standard errors of the estimated coefficients.

To deal with these three issues simultaneously, we use a village random effects tobit model, which assumes that the unobservable for each village is drawn from the same normal distribution. We then adjust the variance-covariance matrix for the presence of the generated regressor using results from Adrian Pagan (1986).[19]

VI. Results of Fertilizer Demand Analysis

A. Sample and variables

Our analysis of fertilizer use is based on the sample of 961 collectively controlled plots from the plot comparison module (see Appendix Table A.1). The dependent variables are, as in Table 1, the quantities of organic and chemical fertilizers per *mu* of plot area. In the case of chemical fertilizer, we use the

[19] The censored least absolute deviations (CLAD) estimator (James L. Powell, 1984), which also deals with problem (i) but does not assume normality, yields similar findings. However, since the CLAD estimator, not being MLE, cannot be easily adjusted for generated regressors, we do not report these results in this paper. Alternatively, we could estimate a conventional tobit model and adjust the variance-covariance matrix for village-level clustering using the Huber-White correction. However, there is no way to correct the standard errors for the generated regressor in this case. Moreover, incorporating the village random effect explicitly into the likelihood function (provided that it is normally distributed), as we do here, yields efficient and consistent parameter and covariance matrix estimates.

amounts applied in the fall growing season (roughly from mid-June to October), where maize, the principal crop, is grown on 78 percent of sample plots. For organic fertilizer, however, we use the total amount applied for the entire cropping year, including the summer growing season (roughly from November of the previous year until early June). During the summer season, wheat is grown in some of the villages almost exclusively on irrigated land (348 plots). Since farmers usually undertake just one manure spreading operation per year, ignoring the summer application would seriously understate soil quality investment on irrigated plots.

Among the theoretical arguments of the fertilizer demand functions, $(\theta, r, \delta, c_x, c_n, \varphi)$, are three—the discount rate, r, the depreciation rate, δ, and the marginal cost of organic fertilizer, c_x—that we do not observe and must therefore treat as fixed parameters. Organic fertilizer is not a purchased input and, to the extent that there is significant variation in its unit cost across households, this must go into the error term. However, later we consider some plausible proxy variables for c_x. For c_n, the marginal cost of chemical fertilizer, we calculate village average prices for pure nitrogen fertilizer based on the household-level data. Given the dearth of observations within each village, we do not include a price for phosphate fertilizer. In any case, the two prices would likely be highly correlated given that most price variation in the cross section is due to transport costs differences. We assume that all other inputs can be freely purchased or rented and that their prices do not vary across the households or villages in our sample, though we relax the first assumption later.

The theoretical model ignores crop choice. Crop rotation, a practice that can enhance soil quality, probably accounts for most of the 209 plots that do not grow maize in the fall season. Since the marginal product of a given fertilizer may differ across crops, it is tempting to include crop controls in the fertilizer regressions. We resist this temptation, first, because crop choice may be endogenous (i.e., correlated with unobserved soil characteristics that also influence fertilizer use) and, second, because crop choice may respond to the degree of expropriation risk. For example, when expropriation risk is high, and therefore the shadow price of organic fertilizer is high, farmers may avoid rotating into crops that are responsive to organic fertilizer. In this case, crop controls will remove part of the impact of expropriation risk on investment that we seek to measure.[20] We provide some evidence on this issue below.

[20] It is important, however, to control for crop prices in the fertilizer demand equations, but we find that they do not vary in the cross-sectional sample for all practical purposes.

Plot fertility, φ, is partly captured by plot-level characteristics, specifically land quality, topography, and irrigation, the same variables used in the hazard analysis. We include a village-level indicator of organic matter in the soil (measurements were taken at six randomly selected sites in each village and averaged) as well. Lastly, we control for tenure type—i.e., for whether the plot is responsibility, ration, or contract land—since tenure type may capture other aspects of land rights besides differences in expropriation risk.[21]

Before turning to the results, it is worth reiterating that our identification of expropriation risk effects relies almost exclusively on *between* village variation in land management policy. Indeed, a regression of $\log\hat{\theta}_i$, predicted for each of the 961 plots in our sample, on the 30 village dummies has an R^2 of 0.95, which means that practically all the variation in the predicted expropriation hazard is across villages. Plot characteristics explain most of the remaining variation in $\log\hat{\theta}_i$. Since the plot variables are also included in the fertilizer regressions, however, they do not contribute to the identification of the expropriation risk effects.

B. Baseline Estimates

Table 4 presents the demand function estimates for organic, nitrogen, and phosphate fertilizer. The results confirm the key prediction of our theoretical model: Higher expropriation risk decreases organic fertilizer use. In contrast, the estimated expropriation risk effects for the two chemical fertilizers do not approach statistical significance.[22] Indeed, the estimates in Table 4 show a remarkable similarity to the private plot dummy coefficients in Table 1, a point to which we return later. Recall from section II that the theoretical predictions for the cross-price effects of expropriation risk on chemical fertilizer use are ambiguous; they depend on the substitution relations among all the inputs. The estimates show, however, that higher expropriation risk does not necessarily decrease the use of *all* fertilizers, only the fertilizer

[21] We do not control for plot size because the dependent variables are in per *mu* terms already. If there is measurement error in plot size, introducing it as a regressor would result in "division bias", leading to a spurious negative relationship between fertilizer use per *mu* and plot size.

[22] The hypothesis of no village random effects is strongly rejected across specifications and, as expected, allowing for village random effects dramatically raises standard errors in some cases. For example, the random effect tobit standard error on the $\log\hat{\theta}_i$ coefficient for organic fertilizer (before adjusting for the generated regressor) is about 2.5 times larger than that of a conventional tobit. By contrast, the correction for the generated regressor raises this standard error by only 8 percent (though by slightly more in some of the other specifications reported below).

with an investment component. Thus, expropriation risk does not act like a tax on output that reduces input intensity across the board; rather, it acts like a tax on investment. Another way to think about the results in Table 4 is that they rule out the possibility that farmers in villages with high expropriation risk have certain unobserved characteristics that mitigate input use overall. If village level unobservables correlated with expropriation risk are present in the data, they must be affecting the marginal products of chemical and organic fertilizers differently. We revisit the issue of village level unobservables shortly.

Besides expropriation risk, practically the only other statistically significant determinant of organic and inorganic fertilizer use is irrigation. Irrigated plots receive considerably less organic fertilizer but more nitrogen than nonirrigated plots (Table 4, row 7).[23] Subject to the caveats discussed above, we also present in Appendix Table A.2, fertilizer demand estimates that account for the fall crops grown on each plot (maize, rice, cotton, or soybeans), and, in the case of organic fertilizer (which is an annual variable), for wheat, the summer crop. Although fertilizer use patterns do differ significantly across certain crops, the coefficients on $\log \hat{\theta}_i$ are more or less the same as in Table 4. Thus, it does not appear that expropriation risk drives cropping patterns. We next consider whether our main finding, the negative relationship between organic fertilizer use and expropriation risk, can be explained by the omission of other important variables from the baseline model.

C. Robustness Checks

Under the assumption of complete factor markets, fertilizer demand does not depend on household input endowments. We now relax this assumption by letting organic fertilizer use per *mu* depend on three farm asset stock variables (normalized by total landholdings): value of farm machinery, value of draft animals, and the number of pigs. The latter variable may be important for organic fertilizer use because pigs, like draft animals, are often a source of manure in China. Specification (1) of Table 5 shows that all three of the assets are positively and significantly associated with organic fertilizer use, but the results for expropriation risk are not appreciably affected.

Specification (2) shows a similar test for the separability of organic fertilizer use with respect to the household's labor endowment, with similar results. In this case, we include three proxies for the shadow

[23] The positive relationship between irrigation and nitrogen use is a common finding in Asian agriculture (e.g., Prabhu L. Pingali, et al., 1997) and derives from the fact that water facilitates the uptake of nitrogen. The result for organic fertilizer might be explained by the fact that this fertilizer improves the moisture retention capacity of the soil. Farmers may thus apply more organic fertilizer to nonirrigated plots to better exploit water from rainfall.

price of household labor: the number of adult males per *mu* of total landholdings, the average years of schooling of adult males, and the presence of a village-owned enterprise. These variables have the expected effects on organic fertilizer, in that households with more manpower apply more and those with better off-farm employment opportunities apply less. Evidently, however, these variables are not strongly correlated with expropriation risk.

Credit constraints or lack of consumption insurance may lead to a different misspecification of the fertilizer demand functions. In either case, production and consumption decisions would be nonseparable and fertilizer use would no longer be independent of wealth. One explanation for our findings might then be that wealthier households happen to live in villages with more frequent land reallocations. At the same time, wealthier households, being less cash constrained, purchase more chemical fertilizers, and therefore use organic fertilizer less intensively, assuming that chemical and organic fertilizers are substitutes. To assess the relevance of this kind of explanation, we control for household wealth using the log of per capita household expenditures as a proxy. Under the null hypothesis of separability, organic fertilizer use should not depend on per capita expenditures. Although the results shown in specification (3) of Table 5 reject this null hypothesis, controlling for household wealth has little effect on the $\log\hat{\theta}_i$ coefficient. This is true despite the fact that village mean expenditures is indeed significantly positively correlated with village mean expropriation risk, so that wealthier villages *do* have more frequent land reallocations.

Next, we address the endogeneity issue that arises from a correlation between $\log\hat{\theta}_i$ and unobserved average village soil fertility, infrastructure, or other productive endowments that may affect the return on organic fertilizer use. To deal with this problem, we construct a direct measure of the village endowment, denoted by $\overline{\varphi}_v$, from the maize yield production function estimated below in Section VII (see Table 6). Computing the production function residuals for all 859 available maize plots (see Appendix Table A.1), and taking village means, gives an estimate of $\overline{\varphi}_v$, albeit a potentially noisy one. To obtain a more accurate estimate of $\overline{\varphi}_v$, we drop 7 villages with fewer than 10 sampled maize plots, paring our fertilizer demand estimation sample down to 853 plots in 24 villages.[24] As shown in specification (4) of Table 5, including this estimate of $\overline{\varphi}_v$ hardly affects the expropriation risk

[24] To avoid spurious correlation between the village average residual and organic fertilizer use on a particular plot, we use leave-out means to calculate $\overline{\varphi}_v$, though this barely makes a difference in practice.

coefficient.[25] Of course, one explanation for this finding is that our estimate of $\overline{\varphi}_v$ is simply too noisy, but this argument is belied by the fact that it attracts a strongly significant coefficient.[26]

Our final specification check concerns whether farmers who report knowing when their contract expires on a particular plot behave differently on that plot. Recall that knowing one's contract expiration date (assuming farmers view commitments about future reallocations by their village leaders as credible) is tantamount to a postponement of tenure insecurity. If so, farmers should not only invest more on these more secure plots, *ceteris paribus,* they should also be less responsive to current expropriation risk. However, to avoid including farmers who may have only found out that their contract had expired during the year of the survey (and, hence, who actually made their investment decisions without knowing when their plots were going to be reallocated), we focus on the 164 plots for which the contract expiration date was known and reported to be after 1995. We also allow for an even more restrictive definition of "secure" plots by distinguishing the 89 plots for which the contract expiration was reported to be in the year 2000 or later. Specifications (5) and (6) in Table 5 show the results of the organic fertilizer tobit including dummy variables corresponding to these categorizations, as well as interactions between these dummies and $\log\hat{\theta}_i$. Only in the case of the more restrictive definition of secure plot in specification (6) do we find effects that approach significance. In particular, on plots that village leaders have assured farmers that they will be able to hold on to for at least the next five years, more organic fertilizer is applied and the responsiveness of organic fertilizer to current expropriation risk is attenuated compared to other plots. Even so, we are not talking about many plots (9 percent); expropriation risk is clearly a salient issue for farmers of the vast majority of plots in our sample.

D. Collectively Controlled versus Private Plots

In light of these results, what can be said about the finding reported in Table 1 that organic fertilizer is used more intensively on private plots than on collectively controlled plots? Recall that while expropriation risk is essentially zero on private plots, other dimensions of property rights may also be

[25] On a technical note, the quadrature method used for numerical integration of the village random effects breaks down with the reduced sample, so we use simulated maximum likelihood instead (taking 100 draws from the normal distribution to simulate the random effects).

[26] That this coefficient is negative indicates that, like irrigation, the village endowment (or, more precisely, the set of productive factors it represents) is a substitute for organic fertilizer. However, at the village level, this endowment is not significantly correlated with expropriation risk, so it does not appear that village land management policy is influenced by these factors.

stronger on these plots than on collectively controlled plots. For this reason, our evidence in Section I for the tenure security hypothesis was only circumstantial. To assess the relative importance of expropriation risk, we can compare the tobit coefficient in Table 1 with that in Table 4. The latter coefficient tells us by how much latent organic fertilizer demand rises if expropriation risk is reduced by 100 percent, i.e., to zero. The answer is 1.35 cubic meters per *mu*. This number is very close to the predicted difference in latent organic fertilizer demand across private and collectively controlled plots from Table 1, which is 1.40 cubic meters per *mu*. It would appear, therefore, that differences in expropriation risk, not other aspects of property rights, explain the bulk of the difference in investment across private and collective plots.

VII. Social Gains from a Tenure Guarantee

There is considerable debate about the urgency of land rights reform in China, with those advocating greater security of ownership pointing to the need to stimulate farm investment. One policy option would be to truly privatize collectively controlled plots, although this appears to be politically unfeasible today. Another option would be for the central government to guarantee plot tenure for a given length of time, after which the plot could again be reallocated according to the discretion of local leaders. Something like this policy has been proposed in China at various times (see Kung and Liu, 1997, for details of recent debates). In this section, we use our model to predict the social gains from such a policy, focusing on the potential impact on maize yield.

What would be the effect of guaranteeing tenure for, say, t_g years in the context of our model? Assuming that the central government is credible and can exert sufficient control over village leaders, expropriation risk will be zero for all $t < t_g$ and will rise to θ for $t \geq t_g$. We cannot use equation (11) to calculate the social gains from this policy because this formula is predicated on a *permanent* change in θ and also assumes zero duration dependence, which we now know to be counterfactual. Thus, return to equation (9), and observe that the social value of the plot after the policy change is

$$(16) \qquad V' = \int_0^{t_g} e^{-rt}\left[R^*(t) - \frac{\partial R^*(t)}{\partial \theta}\theta\right]dt + e^{-rt_g}\int_0^{\infty} e^{-rt}\{S(t)R^*(t) + f(t)V'\}dt$$

The first term in equation (16) reflects the fact that in the t_g years during which θ is zero the farmer receives a higher net yield ($\partial R^*(t)/\partial\theta < 0$) and receives it with certainty.[27] We assume that the value function is approximately stationary both in equation (9) and in equation (16). Thus, the social value of the plot prior to the policy change, *V*, is given by equation (10). We wish to calculate the percentage difference $\Delta V/V = (V' - V)/V$.[28]

To implement this calculation, we must make some assumptions about the net yield function, $R^*(t)$. First of all, as in our empirical specification, we assume that organic fertilizer use, $x^*(t) = x^*$, is constant over time. $R^*(t)$ depends on the marginal cost of organic fertilizer, c_x, which is unobserved, but we can use the first-order condition (4), assuming an interior solution, to eliminate it. Also, the cost of other variable inputs, such as labor and animal power, not to mention chemical fertilizers, should be deducted in calculating $R^*(t)$. If we assume that all of these non-investment inputs are used to the point where price equals marginal product, and that the production function is locally linear, then we have $R^*(t) = y_0 + q(t)F_x x^*$, where $q(t) = (r + h(t))/(r + \delta + h(t))$ and y_0 is expected yield net of the productive contribution of all variable inputs.[29] Further, we can use the fact that $(\partial R^*(t)/\partial\theta)\theta = [c_x(r + h(t))/\delta](\partial x^*/\partial\theta)\theta \; = q(t)F_x\,(\partial x^*/\partial\theta)\theta$. Thus, to calculate $\Delta V/V$ we need to know the discount rate *r*, the depreciation rate δ, the hazard function $h(t)$, the marginal product of organic fertilizer F_x, y_0, and $(\partial x^*/\partial\theta)\theta$.

Neither *r* nor δ are known, but we can take an educated guess at their values. Since the effects of organic fertilizer are said to last about four or five years, if we assume that a unit of organic fertilizer is depleted by ninety percent in four years, we get $\delta = 0.58$ by solving $e^{-\delta 4} = 0.1$. We also use values of δ corresponding to 90 percent depletion after 3 years ($\delta = 0.77$) and after 5 years ($\delta = 0.46$).

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[27] Actually, this is only an approximation because as plot tenure approaches t_g the farmer will have an incentive to deplete soil quality toward its optimal level under tenure uncertainty. However, as long as we consider values of t_g in the relatively distant future, the approximation should be reasonably good.

[28] In the absence of duration dependence, $\lim_{t_g \to \infty} (V' - V)/V$ is precisely η_θ as defined in equation (11).

[29] So yield $= y_0 + F_n n^* + F_x x^*$ and cost per *mu* $= F_n n^* + \delta F_x x^*/(r + \delta + h(t))$, in the two-input case.

The marginal product of organic fertilizer, F_x, is estimated from a yield production function using only the plots on which maize is grown.[30] Table 6 presents household fixed-effects estimates of the production function, using 314 households with two maize plots in the plot comparison module.[31] The linear specification may be viewed as a first-order approximation to the true production function, and should estimate the *average* marginal product in the sample correctly. Linearity is necessary to deal with the heavy censoring of fertilizer use, and other variables, at zero. It would in any case be difficult to estimate higher order terms with any precision given the fixed effects procedure. The advantage of fixed effects is that it removes household level (but not plot level) unobservables that are likely to be correlated with input choices. One issue that remains, however, is that certain inputs might be partly chosen after the plot-specific production shock is revealed and so may be endogenous. To assess this problem, we include a subjective measure of the output shock (percentage output shortfall on each plot reported by the farmer in the survey) in specification (2) of Table 6. This shock variable attracts a significantly negative coefficient, but hardly affects the variable input coefficients.[32] The estimate of the marginal product of organic fertilizer, $\hat{F}_x$, is about 18 kg maize per cubic meter, which at the production function sample means corresponds to a yield elasticity of around 6 percent. We also use the estimates in Table 6 to calculate y_0 for all 961 plots in our fertilizer sample based on their fixed characteristics. We include all collectively controlled plots in our calculation, not just those on which maize is grown, because all of these plots could *potentially* grow maize.

For $(\partial x^* / \partial\theta)\theta$ we use the marginal effects based on the tobit estimate in Table 4. The average marginal effect over all 961 plots is -0.61, which is considerably lower than the coefficient estimate of -1.35 due to the heavy censoring; it corresponds to an ε_θ of -0.374 at the sample mean. All the integrals

[30] F is actually a function of the stock k, but we make the substitution $k \cong x/\delta$ (see equation (6)). Also, recall that we use the annual amount of organic fertilizer (summer + fall season) in the demand analysis, and we do likewise for the production function estimation. However, to correct for intra-year depreciation we discount fertilizer applied in the previous summer season using $\delta = 0.58$. Our production function estimates are robust to the choice of δ in the range used for Table 7.

[31] The fact that more intensive maize cultivators might be a select sample of farmers should not bias the estimates of the maize production function because the fixed effects procedure purges all household level unobservables. Irrigation hardly varies across plots within the same farm; there are only ten dual-maize plot households with one irrigated and one non-irrigated plot. To avoid basing our estimates on minimal within-farm variation, we drop the plots of these 10 households from the production function sample and delete irrigation from our list of plot characteristics. For the remaining plots (the 314 used in Table 6), the effect of irrigation on yields is impounded in the household fixed effect.

[32] One caveat is that the shock variable only records the magnitude of negative shocks, not positive ones.

in the expressions for V (equation (10)) and V' (equation (16)) are evaluated numerically. Table 7 reports the *median* value of $100 \text{x} \Delta V/V$ across the 961 collectively controlled plots at different values of r and δ, and for a t_g of 15 and 30 years. In all cases, the efficiency gains from a tenure guarantee are minimal. The median increase in plot value never exceeds one percent, regardless of the configuration of parameter values. Despite the large estimated elasticity of investment with respect to expropriation risk, ε_θ, the gains from the tenure guarantee are limited by the relatively small productive contribution of organic fertilizer, as indicated by its low yield elasticity.[33] As for the effect of a longer period of guaranteed tenure, this is proportionally greater at lower discount rates; when farmers discount the future heavily, there is essentially no difference in social gains between a 15 and 30 year tenure guarantee.

Our finding that the efficiency costs of tenure insecurity are small rests, of course, on several simplifying assumptions, as well as on empirical estimates that are subject to uncertainty. Regarding the latter, one concern might be that, because of attenuation bias due to measurement error in organic fertilizer use, our estimate of the marginal product of organic fertilizer from Table 6 is too low. However, even if as much as one half of the variance in organic fertilizer use within households is due to random measurement error – so that the true marginal product is actually about twice as large as our estimate -- the percentages in Table 7 would only double. This is not enough to change our basic conclusion. The effect of unobserved plot level characteristics (omitted variable bias) on our marginal product estimate is more difficult to assess. Still, to obtain appreciable efficiency costs in our welfare analysis, the true marginal product of organic fertilizer would have to be at least one order of magnitude larger than our estimate, which seems unlikely.[34]

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[33] Our calculations do not take into account the impact of organic fertilizer use on summer wheat yields. However, to the extent that organic fertilizer acts as a way to provide moisture to crops, it would be less important in wheat production, which is almost all irrigated, than it is for maize, much of which is rainfed. This is confirmed when we estimate a wheat yield production function analogous to that in Table 6 using a sample of 121 farmers that cultivate two wheat plots. The coefficient on organic fertilizer in this case is actually negative, but insignificant. If we take the marginal product of organic fertilizer in wheat production to be essentially zero, then the cultivation of this summer crop only affects the social gain calculations by raising V for these 348 plots; this would make the numbers in Table 7 even lower.

[34] For example, a referee suggests that the previous year's organic fertilizer use is one such omitted plot characteristic, and that it might be *negatively* correlated with current use if manure is spread only every other year. We can get an upper bound on the coefficient bias in this scenario by assuming (1) current and lagged organic fertilizer use are the only variables in the production function regression, (2) they have equal effects on yields, (3) they have equal variances, and (4) they are perfectly negatively

A more fundamental assumption underlying our efficiency calculations is that all the benefit from reduced expropriation risk comes through an increase in organic fertilizer use. In section I, we discounted the impact of plot-level expropriation risk on fixed investments in land. Yet, there may be important *recurrent* investments in land that our analysis ignores. In the rest of this section, we examine two such activities: plot maintenance and crop rotation.

Our data set provides information at the household level on how family members allocated their time among several agricultural tasks.[35] For instance, the survey asks about the number of days devoted to organic fertilizer application and to land maintenance. This latter category includes land leveling, cleaning and preparing irrigation and drainage ditches, and bunding (it excludes plowing and furrowing). At least some of these activities may have effects that carry over to future years. Land maintenance accounts for around 6 percent of total crop production time compared to 8 percent for organic fertilizer application.

How does the intensity of land maintenance effort respond to expropriation risk? Table 8 addresses this question by regressing effort intensity on predicted expropriation risk as well as on the other plot and village level variables used in our baseline fertilizer specifications in Table 4.[36] To aggregate plot-level variables, including $\log \hat{\theta}_i$, to the household level, we take household weighted averages using the ratio of plot area to total household land area as weights.[37] To deal with the extensive censoring of the dependent variables at zero and the village clustering, we again use the tobit model with village random effects. Specification (1) indicates that higher expropriation risk significantly reduces time spent applying organic fertilizer, a finding that corroborates our earlier results.[38] Specification (2), however,

correlated. In this case the true marginal product would only be twice as large as the estimated marginal product.

[35] We do not have information on hired labor disaggregated by task. However, in our sample, hired and exchange labor account for only about 8 percent of total annual labor days in farm production. Moreover, most hired labor in rural China is used in harvesting and weeding, rather than in land maintenance, so ignoring it should not affect the present analysis.

[36] All except distance of plot from the house, which is only available for plots included in the plot comparison module.

[37] In the case of private plots, which constitute only 3.4 percent of land area for the average household in our sample, we set $\log \hat{\theta}_i$ to a small number. The choice of this number is inconsequential because we also control separately for the fraction of private land in the regressions.

[38] We do not correct the standard errors in Table 8 for the generated regressor (as we did before), because the aggregation of plot level predicted hazards to the household level greatly complicates the procedure.

shows no significant impact of expropriation risk on land maintenance effort. One interpretation of this result is that land maintenance effort is unresponsive to tenure insecurity. Alternatively, it could be that most plot maintenance activities simply do not have long-lasting effects and, instead, must be repeated annually to prepare fields for cultivation.

Turning next to crop rotation, we mentioned in Section VI that this practice can enhance soil quality and may be viewed as an investment to the extent that farmers forgo current revenue by temporarily growing a less lucrative crop. The plot comparison module of our survey asks farmers which crop they grew the previous year (fall, 1993), as well as which crop they were planning to plant the following year (fall, 1995). We can use this information to construct plot-level indicator variables for whether crops were rotated in 1993-94 and for whether they will be rotated in 1994-95. Missing values reduce our samples (from 961) to 867 plots for 1993-94 and to 891 plots for 1994-95.[39] Of these total plots, farmers rotated 77 (9 percent) in 1993-94 and 70 (8 percent) in 1994-95. Specifications (3) and (4) in Table 8 presents village random effects probit models, analogous to our baseline fertilizer specifications, for the two crop rotation variables. In neither case does the effect of expropriation risk come close to statistical significance. A likely explanation for this result, and for the low incidence of crop rotation in our sample, is that farmers find it cheaper to use organic fertilizer for rebuilding soil structure than to rotate their crops. However, farmers occasionally do switch crops to exploit potential profit opportunities; indeed, most rotations are among different cash crops (e.g., maize-cotton, maize-soybeans) whose relative prices may change from year to year.

To sum up, in contrast to the strong and robust finding for organic fertilizer use, it does not appear that expropriation risk influences other plot-specific recurrent investments. Although it is possible that plot maintenance effort and crop rotation are not measured as well in our data set as is organic fertilizer use, or that some other plot-specific investments were entirely missed by the survey instrument, the evidence that we can report suggests it is reasonable to focus solely on the benefits of greater organic fertilizer use in our policy simulations.

As noted above, however, in Tables 4 and 5 this correction raises standard errors by no more than around ten percent, which would not be enough to affect inferences in Table 8.

[39] Most of the missing values for the 1993-94 rotation variable are due to the fact that farmers who were allocated plots very recently have not had the opportunity to grow a crop in 1993. Because of this, the sample may be selected in a way that is correlated with expropriation risk. For the 1994-95 rotation variables, missing values may arise either because farmers have not yet decided what crop they will grow, or because they already know they will lose the plot in 1995.

VIII. Conclusion

In this paper, we examine the risk of land expropriation as a constraint on farm investment in rural China. We have argued that, perhaps more readily than other aspects of property rights, expropriation risk can be quantified. Plot tenure data provide a mirror that reflects the recent history of land expropriation. Using a hazard model, we have been able to extract the exogenous risk of expropriation from such data. Having this objective measure of tenure insecurity allows us to estimate a structural model, which makes possible an assessment of the social benefits of policies designed to reduce tenure insecurity.

Our empirical results strongly support the view that heightened expropriation risk puts a damper on investment in rural China. Expropriation risk varies considerably across the 31 villages in our sample due to differences in local land management policy. Importantly, farmers living in villages where expropriation risk is higher use organic fertilizer less intensively. This is not the case with chemical fertilizers, which are known to have no long-lasting effects on soil quality, nor is it the case with plot maintenance effort or crop rotation, which could have an investment component. Despite having a significantly negative impact on one form of plot-specific investment—organic fertilizer use—periodic land reallocations do not appear to entail a substantial social cost, at least in this part of China. According to our estimates, organic fertilizer is simply not an important enough input for underutilization to matter much. Meanwhile, many of the more capital-intensive agricultural investments, such as canal irrigation, drainage, and terracing projects, are being (or have already been) undertaken at the village level. Other fixed investments, such as wells, are less dependent on plot-specific rights. And, while it is likely that tenure insecurity would affect a farmer's willingness to invest in orchards or other activities that require a relatively permanent transformation of cropland, the social cost of this distortion is likely to be small since it would, at most, affect only a small fraction of China's arable land.

Of course, the key question for a complete evaluation of China's land management system is the extent to which the investment decisions of village leaders are efficient. We have argued here that due to the public goods and externalities involved in these investments some degree of delegation to local authorities on the part of individual farmers would be efficient, and this should be true for rural communities outside of China as well. Perhaps, however, in response to the poor investment climate induced by land reallocations, village leaders in China have taken over more decision-making power from individual farmers than would otherwise be efficient. Perhaps, too, village leaders in China mismanage these investments more so than do communal institutions in other countries. These are important and interesting questions for future research and must qualify our conclusions about the social costs of tenure insecurity in China.

Appendix A

Derivation of the First-Order Conditions

The Hamiltonian for the problem is

(A.1) $$\mathscr{H} = e^{-rt}S(t)R(t) + \lambda(t)[-\delta k(t) + x(t)] + \nu_1(t)x(t) + \nu_2(t)n(t)$$

where $\lambda(t)$, $\nu_1(t)$, and $\nu_2(t)$ are multiplier functions, with $\nu_1(t) > 0$ as $x(t) = 0$, $\nu_1(t) = 0$ otherwise, and $\nu_2(t) > 0$ as $n(t) = 0$, $\nu_2(t) = 0$ otherwise. The first-order necessary conditions for a maximum are

(A.2) $$-e^{-rt}S(t)c_x + \lambda(t) + \nu_1(t) = 0$$

(A.3) $$e^{-rt}S(t)\left[F_n(t) - c_n\right] + \nu_2(t) = 0$$

(A.4) $$-\dot{\lambda}(t) = e^{-rt}S(t)F_k(t) - \delta\lambda(t)$$

To get equations (4) to (6) in the text, set $\nu_1(t) = \nu_2(t) = 0$ and differentiate (A.2) with respect to *t*. Combining the result with (A.4) yields equation (4). Equation (5) is immediate from (A.3). Differentiating (4) and (5) with respect to *t* and combining with constraint (2) yields equation (6).

Appendix B

Likelihood Function for the Hazard Model

Following Heckman and Singer (1984), we allow for multiplicative heterogeneity of the form $\mu_i\theta_i$ and assume a discrete distribution for μ_i with *M* points of support $\{\mu_1, \ldots, \mu_M\}$, each occurring with probability $\pi_m, m = 1,2,\ldots, M$. The distribution of completed tenure for plot i, unconditional on μ_i, is then $f(t_i;\theta_i) = \sum_{m=1}^{M}\pi_m f(t_i;\mu_m\theta_i)$. Using equation (13), it follows that the distribution of incomplete plot tenure is[40]

[40] Steven W. Salant (1977) considers such a likelihood with a parametric heterogeneity distribution. See also Mark Gersovitz, et. al (1998) for a nonparametric application.

(B.1) $$f^*(t_i^*;\theta_i) = \frac{\sum_{m=1}^{M}\pi_m S(t_i^*;\mu_m\theta_i)}{\sum_{m=1}^{M}\pi_m \int_0^\infty S(\tau;\mu_m\theta_i)d\tau}$$

The full likelihood is the product of $L_i = f^*(t_i^*;\theta_i)^{1-d_i}[1-\Pr(t_i^* < t_R)]^{d_i}$ over all plots, where d_i is the left-censoring indicator. In the case of the Weibull hazard, the likelihood contribution of plot i becomes

(B.2)

$$L_i = \left[\frac{\theta_i^{1/\alpha}}{\omega}\sum_{m=1}^{M}\pi_m \exp\{-\mu_m\theta_i(t_i^*)^\alpha\}\right]^{(1-d_i)} \times \left[1-\frac{1}{\omega}\sum_{m=1}^{M}\pi_m\Gamma(1/\alpha,\mu_m\theta_i(t_R)^\alpha)\mu_m^{-1/\alpha}\right]^{d_i}$$

where $\omega = \Gamma(1+1/\alpha)\sum_{m=1}^{M}\pi_m\mu_m^{-1/\alpha}$, $\Gamma(x)$ is the gamma function, and $\Gamma(x,y)$ is the incomplete gamma function. While equation (B.2) does not, strictly speaking, nest the model with no unobserved heterogeneity, we can conclude such heterogeneity is not important if the data support only $M = 1$; this is the case if either π_1 or π_2 is estimated to be very close to zero when $M = 2$, or if μ_1 is very close to μ_2. We continue to add points of support beyond $M = 2$ until the data fail to support doing so according to these criteria.

Finally, note that this likelihood function allows only plot-level heterogeneity, not heterogeneity in the form of household or village level random effects, because the random effect must be integrated out of both numerator and denominator of (B.1), the plot-level likelihood contribution.

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Table 1

Fertilizer Use on Private versus Collective Plots: Fixed Effect Tobit Estimates

	Organic cubic meters/*mu* (1)	**Nitrogen** kilograms/*mu* (2)	**Phosphate** kilograms/*mu* (3)
Fixed effect tobit coefficients:			
Private plot	1.40	-0.77	0.67
	[4.00]	[0.70]	[0.55]
High quality	0.07	1.93	-1.29
	[0.24]	[0.91]	[0.66]
Flat topography	-0.83	3.68	0.97
	[1.64]	[1.55]	[1.28]
Distance from house (km.)	-0.24	-0.99	0.02
	[0.43]	[0.74]	[0.03]
Means (standard deviations):			
Private plots	4.61	16.3	1.95
	(3.38)	(10.6)	(3.27)
Collective plots	3.46	16.0	1.77
	(2.94)	(8.95)	(1.86)
Difference	1.14	0.29	0.19
	(2.88)	(8.58)	(3.59)
Percent observations censored	12	8	54

Notes.— Absolute t-values in square brackets. 15 *mu* = 1 hectare. Sample size for all regressions is 216 plots (108 households). Fixed effect tobit uses quadratic loss function (see Honoré, 1992).

Table 2

Frequency of Plot Tenure Durations: By Type of Collective Land

Duration (years)	Responsibility	Ration	Contract	All Types
1	327	35	153	515
2	254	43	47	344
3	146	25	39	210
4	305	40	26	371
5	105	33	49	187
6	107	31	34	172
7	33	3	5	41
8	28	4	4	36
9	72	13	0	85
10	115	1	13	129
11	67	6	8	81
12	199	17	16	232
13	310	16	11	337
14	83	1	0	84
15	25	0	3	28
16	39	3	4	46
Total	2215	271	412	2898
Mean	6.88	5.27	4.00	6.32
(Standard Deviation)	(4.80)	(3.80)	(3.70)	(4.69)

Notes.—Eight durations are truncated at 16 years.

Table 3
Hazard Analysis of Tenure on Collective Plots

	Mean (Standard Deviation)	**Unobserved Heterogeneity Control** *None* **(1)** [a]	*Nonparametric* **(2)** [a, b]	**Village Fixed Effects** **(3)** [c, d]	**(4)** [c]
Duration dependence		0.034	1.92	0.335	0.325
($\log \alpha$)		(0.07)	(2.44)	(5.30)	(5.19)
Household characteristics:					
Age of head	44.3	-0.021	-0.021	-0.013	-0.013
	(11.7)	(4.58)	(2.38)	(3.03)	(3.40)
Size	3.81	-0.098	-0.032	-0.082	
	(1.11)	(2.26)	(0.67)	(2.09)	---
Adult males per	0.516	-0.177	-0.159	-0.206	
capita	(0.160)	(6.45)	(0.63)	(0.86)	---
Children age 0-15	0.229	-0.645	-0.881	0.017	
per capita	(0.198)	(3.54)	(1.97)	(0.08)	---
Plot characteristics:					
Ration plot	0.094	0.365	0.306	-0.165	-0.168
		(2.68)	(0.89)	(1.54)	(1.52)
Contract plot	0.142	0.694	0.704	0.405	0.401
		(2.65)	(4.82)	(4.02)	(3.99)
Area (*mu*)	2.74	0.026	0.028	0.0046	0.0036
	(4.78)	(4.23)	(9.92)	(1.79)	(1.43)
High quality	0.321	-0.041	-0.058	-0.0024	-0.0082
		(0.52)	(0.62)	(0.04)	(0.12)
Irrigated	0.414	-0.053	-0.118	-0.068	-0.078
		(0.13)	(0.52)	(0.54)	(0.61)
Flat topography	0.802	0.474	0.403	0.245	0.229
		(15.1)	(1.53)	(2.51)	(2.29)
-Log-likelihood		6,488	6,418	5,917	5,926

Notes.— Sample size is 2,898 plots (719 households). All models include a constant term.

[a]Absolute t-values (in parentheses) based on robust covariance matrix adjusted for village-level clustering.

[b]Estimates of the heterogeneity distribution parameters (robust t-values in parentheses): μ_1 normalized to zero and $\hat{\pi}_1 = 0.629 (4.50)$; $\hat{\mu}_2 = -2.19\ (4.94)$ and $\hat{\pi}_2 = 0.063\ (1.48)$; $\hat{\mu}_3 = -1.09\ (4.23)$ and $\hat{\pi}_3 = 0.307\ (1.69)$.

[c]Absolute t-values (in parentheses) based on robust covariance matrix adjusted for household-level clustering.

[d]Wald test (based on robust covariance matrix) for H_0: no village fixed effects, $\chi^2_{(30)} = 486$ (p-value<0.00001); for H_0: no household demographic effects, $\chi^2_{(3)} = 4.92$ (p-value=0.18).

Table 4
Fertilizer Use On Collective Plots: Village Random Effects Tobit Estimates

	Mean (Standard Deviation)	**Organic** cubic meters/*mu* **(1)**	**Nitrogen** kilogram/*mu* **(2)**	**Phosphate** kilogram/*mu* **(3)**
$\log \hat{\theta}_i$	-3.27	-1.35	-0.344	0.380
	(1.21)	(3.56)	(0.42)	(0.52)
Village characteristics:				
Organic matter in soil	1.20	0.160	-2.87	-1.01
(percent)	(0.37)	(0.21)	(1.29)	(0.51)
Price of nitrogen	1.65	-0.446	-3.69	1.41
(yuan/kilogram)	(0.17)	(0.15)	(0.78)	(0.17)
Plot characteristics:				
Ration plot	0.171	-0.100	-1.47	-0.037
		(0.17)	(1.05)	(0.03)
Contract plot	0.060	0.545	-1.20	0.706
		(0.76)	(0.53)	(0.54)
High quality	0.421	0.082	1.10	0.531
		(0.23)	(1.03)	(0.72)
Irrigated	0.553	-1.35	3.45	0.203
		(2.85)	(2.63)	(0.24)
Flat topography	0.925	1.43	3.35	0.277
		(2.25)	(1.76)	(0.21)
Distance from house	0.740	-0.628	-0.480	0.148
(kilometers)	(0.638)	(2.40)	(0.62)	(0.34)
Constant		-4.63	24.1	-2.82
		(0.82)	(2.66)	(0.21)
H_0: No village random		43.7	18.6	39.3
effects, $\chi^2_{(1)}$ [p-value]		[0.000]	[0.000]	[0.000]
Mean of dependent variable		1.63	20.8	2.32
(Standard deviation)		(2.82)	(13.4)	(4.26)
Percent observations censored		47	7	55

Notes.— Absolute t-values (in parentheses) based on covariance matrix adjusted for generated regressor. The omitted category for tenure type is responsibility land. Sample size for all regressions is 961 plots (608 households in 31 villages).

Table 5
Alternative Specifications For Organic Fertilizer Use: Village Random Effects Tobit Estimates

	Mean (Standard Deviation)	**Organic Fertilizer** cubic meters/*mu* **(1)**	**(2)**	**(3)**	**(4)[b]**	**(5)**	**(6)**
$\log\hat{\theta}$		-1.27 (3.46)	-1.19 (3.26)	-1.47 (3.39)	-1.29 (3.38)	-1.40 (3.59)	-1.69 (3.47)
Value of farm machinery (yuan/*mu*)[a]	219 (569)	0.0123 (5.29)	---	---	---	---	---
Value of draft animals (yuan/*mu*)[a]	69.1 (158)	0.0334 (3.81)	---	---	---	---	---
Number of pigs per *mu*[a]	0.099 (0.238)	2.07 (3.56)	---	---	---	---	---
Number of adult males per *mu*[a]	0.293 (0.247)	---	1.35 (1.82)	---	---	---	---
Average years schooling of adult males in household	5.66 (2.25)	---	-0.198 (2.94)	---	---	---	---
Village enterprise dummy	0.325	---	-1.88 (1.91)	---	---	---	---
Log total household expenditures per capita	7.63 (0.45)	---	---	-0.845 (2.34)	---	---	---
$\hat{\bar{\varphi}}_v$	20.4 (90.8)	---	---	---	-0.0199 (4.09)	---	---
Contract expires after 1995	0.171	---	---	---	---	1.98 (1.19)	---
Contract expires after 1995 $\times\log\hat{\theta}$		---	---	---	---	0.338 (0.78)	---
Contract expires after 1999	0.093	---	---	---	---	---	5.84 (2.32)
Contract expires after 1995 $\times\log\hat{\theta}$		---	---	---	---	---	1.16 (1.89)

Notes.— Selected coefficients reported. Regressions also include all variables from Table 4 (see notes to Table 4).

[a] Refers to total household landholdings.

[b] Sample size is 853 plots (543 households in 24 villages). Coefficient on $\log\hat{\theta}$ in baseline specification for this sample is –1.26 (2.59)

Table 6

Maize Yield Production Function: Household Fixed Effects Estimates

	Mean (Standard Deviation)	**Maize Yield** kilogram/*mu*[a] **(1)**	**(2)**
Variable inputs:			
Organic fertilizer	2.39	18.9	18.4
(cubic meters /*mu*)	(3.05)	(2.51)	(2.55)
Nitrogen fertilizer	19.4	6.52	6.77
(kilograms/*mu*)	(12.0)	(3.91)	(4.24)
Phosphate fertilizer	2.05	-0.82	-1.02
(kilograms/*mu*)	(4.21)	(0.24)	(0.31)
Labor (days/*mu*)	11.4	7.58	6.86
	(9.95)	(2.09)	(1.97)
Animal power (days/*mu*)	2.05	7.61	5.71
	(3.19)	(0.53)	(0.42)
Plot characteristics:			
High quality plot	0.489	154	129
		(5.61)	(4.86)
Flat topography	0.904	14.0	34.0
		(0.30)	(0.77)
Distance from house	0.654	-9.70	-17.0
(kilometers)	(0.534)	(0.40)	(0.74)
Output shock[b]	11.4		-5.08
	(20.1)	---	(5.44)

Notes.—Absolute t-values in parentheses. Sample size is 628 plots (314 households).

[a]Mean = 709, standard deviation = 347.

[b]Farmers' report of percentage output shortfall relative to normal output.

Table 7

Social Gains from Tenure Guarantee:

Median Percentage Change in Plot Value

	δ	**0.46**	**0.58**	**0.77**
15-year Tenure Guarantee:				
	0.05	0.4	0.3	0.3
r	**0.1**	0.6	0.5	0.4
	0.2	0.9	0.8	0.6
30-year Tenure Guarantee:				
	0.05	0.6	0.5	0.4
r	**0.1**	0.7	0.6	0.5
	0.2	0.9	0.8	0.7

Notes.— See text for details of the calculation of $100 \text{ x } \Delta V / V$.

Table 8
Household Time Use in Crop Production: Village Random Effects Estimates

	Time Allocation Tobit [a]		**Crop Rotation Probit** [b]	
	Organic Fertilizer Application	**Land Maintenance**	**1993-94**	**1994-95**
	(1)	**(2)**	**(3)**	**(4)**
$\log \hat{\theta}_i$	-0.230	0.053	0.071	0.018
	(2.76)	(0.97)	(0.61)	(0.14)
Village characteristics:				
Organic matter in soil (percent)	-0.522	-0.112	-0.136	0.211
	(1.72)	(0.56)	(0.39)	(0.53)
Price of nitrogen (yuan/kilogram)	-0.187	-1.05	1.21	1.34
	(0.30)	(2.87)	(1.42)	(1.42)
Plot characteristics:				
Ration plot	-0.0027	0.020	0.220	0.020
	(0.15)	(1.10)	(1.11)	(0.1)
Contract plot	0.020	0.008	0.107	0.473
	(1.22)	(0.49)	(0.37)	(1.66)
Private plot	-0.069	-0.002		
	(1.31)	(0.03)	---	---
High quality	0.0026	0.003	-0.031	0.044
	(0.28)	(0.29)	(0.18)	(0.23)
Irrigated	-0.0354	0.029	-0.102	-0.014
	(2.60)	(2.12)	(0.52)	(0.07)
Flat topography	-0.013	0.010	-0.627	-0.561
	(0.73)	(0.52)	(2.62)	(2.13)
Distance from house			-0.038	0.038
(kilometers)	---	---	(0.37)	(0.37)
Constant	0.112	0.204	-2.53	-3.67
	(0.92)	(2.79)	(1.64)	(2.1)
H_0: No village random	43.1	17.2	37.2	42.1
effects, $\chi^2_{(1)}$ [p-value]	[0.000]	[0.000]	[0.000]	[0.000]
Sample size	721	721	867	891
Mean of dependent variable	0.0800	0.0598	0.0888	0.0786
(Standard deviation)	(0.0950)	(0.0755)	---	---
Percent observations censored	23	35	---	---

Notes.— Absolute t-values in parentheses.

[a]Time allocation is measured as fraction of annual crop production days by family members. These are household level regressions in which all plot characteristics (including $\log \hat{\theta}_i$) are weighted averages of the corresponding plot-level variable (weights equal to the ratio of the plot's area to total household land area).

[b]Indicator variable takes on value of one if different fall crops were grown on the plot from one year to the next, zero otherwise.

Table A.1 Samples used in the Empirical Analysis

Number of plots (households)	Sample selection criteria	Empirical analysis (Table)
Plot-wise enumeration:		
3,113 (727)	All plots	---
2,898 (719)	Collectively controlled plots	Hazard (2,3)
Plot comparison module:		
1,074 (612)	Valid data/major crops[a]	---
961 (608)	Collectively controlled plots	Fertilizer demand (4,5)
859 (534)	Maize plots of all tenure types	Computing $\hat{\varphi}_v$ (5)
628 (314)	Household has two maize plots	Production function (6)
216 (108)	Household has private and collective plot	Fertilizer use comparison (1)
Household time allocation:		
(721)	Household has nonmissing time-use data	Land maintenance effort (8)

[a]A total of 1256 plots from 662 households are recorded in the plot comparison module. Of these, 182 plots were dropped because of inconsistent responses, missing data on key variables, or because vegetables or other minor crops were grown on these plots.

Table A.2

Fertilizer use on Collective Plots with Crop Controls: Village Random Effects Tobit Estimates

	Mean (Standard Deviation)	**Organic** Cubic meters/*mu* **(1)**	**Nitrogen** kilogram/*mu* **(2)**	**Phosphate** kilogram/*mu* **(3)**
$\log \hat{\theta}_i$		-1.19	-0.857	1.01
		(3.42)	(1.08)	(1.35)
Fall rice	0.135	-1.74	3.09	-0.983
		(2.16)	(1.50)	(0.75)
Fall cotton	0.047	0.407	1.87	3.82
		(0.37)	(0.74)	(2.63)
Fall soybeans	0.035	-0.995	-14.8	-0.686
		(1.06)	(5.62)	(0.49)
Summer wheat	0.362	0.333	---	---
		(0.44)		

Notes.— Selected coefficients reported. Regressions also include all variables from Table 4 (see notes to Table 4). Omitted crop is maize in the fall season and fallow in the summer season.

Figure 1. Predicted Expropriation Risk